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Article (Accepted Version)

Mamatzakis, Emmanuel C and Vu, Anh N (2018) The interplay between quantitative easing, risk and competition: The case of Japanese banking. Financial Markets, Institutions and Instruments, 27 (1). pp. 3-46. ISSN 0963-8008

THE INTERPLAY BETWEEN QUANTITATIVE EASING AND RISK: THE CASE OF THE JAPANESE BANKING

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Abstract

The Japanese banking industry is an interesting one, given chronic problems related to notorious non-performing loans, originated back in the 1990s, but also due to an unprecedented monetary expansion. In this paper, we focus on the impact of quantitative easing on bank level risk, while controlling for bank competition. We opt for a measure of bank specific risk-taking based on a new data set of bankrupt and restructured loans. Given issues related to endogeneity among the main variables, we adopt dynamic panel threshold and panel vector autoregression analyses that address such criticism. Results demonstrate that quantitative easing reduces bankrupt and restructured loan ratios, though we do not observe a similar impact on bank stability. Given the adoption of negative rates in January 2016 by the Bank of Japan, our study comes is timely and provides insightful implications for future research.

JEL-classifications: G21, C23, E52

Keywords: Quantitative easing; bank risk-taking; Japan.

Acknowledgements: This study was conducted under the Bank of Greece's programme of cooperation with universities. The authors are grateful to Heather Gibson for valuable comments and all participants at the research seminar of Bank of Greece. We also thank G. Tavlas, D. Malliaropulos, A. Tagkalakis, G. Hondroyiannis, Thomas Weyman-Jones, Roman Matousek, Manthos Delis, and Efthymios G. Tsionas. Any remaining errors are our own.

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1. Introduction

The association between quantitative easing and bank risk-taking has raised concerns among academics and policymakers alike (Chodorow-Reich, 2014; Claeys & Darvas, 2015). Low short-term interest rates prior to loan issuance result in banks granting more new risky loan portfolios, distorting their credit supply to favour borrowers with worse credit histories, lower ex-ante internal ratings, and weaker ex-post performance (Ioannidou et al., 2015; Jiménez et al., 2014). Lower returns from yields is another motive for financial institutions to accelerate their risk-taking activities (Chodorow-Reich, 2014; Rajan, 2005). Banking surveys based on credit standards in the US and the UK, by contrast, do not suggest an excessive risk-taking by banks as a result of the enforcement of quantitative easing (Claeys & Darvas, 2015). However, the literature has not been settled to date about the impact of quantitative easing on bank risk taking. Herein, we explore a new data set and measure bank risk at bank level using new Japanese nonperforming loans information. In addition, the Bank of Japan has been the pioneer in empowering quantitative easing policies. There has been a strong record of active and aggressive quantitative easing since 2010. We are interested in investigating whether the warning of heightened risks associated with this policy is supported by Japanese bank level data. We hypothesise, based on the aforementioned literature, that quantitative easing could affect bank risk-taking.

After the acute phase of the banking crisis in Japan (1997-1999), the banking system underwent major reforms, bailout and consolidation from 1999 to 2003. As a response to the banking crisis, between 2000 and 2012, quantitative easing was launched twice: firstly, during March 2001-March 2006 and secondly from October 2010 onwards to date. Given this extensive monetary expansion, it is of interest to examine its impact on bank performance. We also control for the degree of competition in the banking industry (Altunbas et al., 2014). We explore in depth the underlying causality among quantitative easing, competition and risk.

Apart from Japan's commitment to do "*whatever it takes*" in achieving growth through quantitative easing, Japan is of interest for the infamous nonperforming loan problem. This destructive effect of the banking crisis prolonged to the first half of the 2000s due to the reluctance of the government in admitting the nonperforming loan issue. Eventually, when doing nothing was too painful to tolerate, the government restructured the whole banking industry. However, this restructuring campaign was not without controversy as it incorporated funding of unprofitable firms, which in turn crowded out solvent firms and lengthened the revitalisation of the economy (Caballero et al., 2006). Moreover, there exists evidence of political influence, with regulators deferring to declare the solvency of banks situated in prefectures supporting the then ruling party (Imai, 2009). When quantitative easing was first introduced, structural reforms of the financial system were essential for the policy to gain its full effectiveness (Bank of Japan, 2001). After the global financial crisis, nonperforming loans of all banks increased slightly from JPY11.4 trillion in March 2008 to JPY12 trillion in March 2009.[1] It is noteworthy that during 2008-2013, the government strategically aimed to assist small and medium enterprises (SMEs), which are Regional Banks' primary corporate clients, via the SME Financing Facilitation Act. One term in the Act involved reclassifying SME's nonperforming loans. This raised the accumulated hidden credit risks within the banking system (Hoshi, 2011), as about 3-6% of total credit in Regional Banks was reclassified (International Monetary Fund, 2012).

Our study contributes to the literature in the following ways. First, we opt for an original data set to capture bank risk that has been overseen by the literature to date. Bank risk-taking, our primary focus, is represented by bankrupt loan ratio and restructured loan ratio. Data on bankrupt and restructured loans are available for Japanese commercial banks and have not been used extensively in the Japanese banking literature (Mamatzakis et al., 2016). We also use the classic measure of bank default risk, the Z-score, to enhance the robustness of our analyses. The use of bankrupt and restructured loans at semi-annual data frequency allows an enriched information set in our modelling of competition and quantitative easing.[2] Second, we employ a bank level proxy of quantitative easing – the bank specific lending rate. The advantages of this microeconomic measure lie in the absence of aggregation bias and the ample set of information. This bank-specific variable ensures its compatibility with the bank level Boone indicator and risks in our analyses. We also conduct the analyses with two other proxies for quantitative easing: the 10-year Japanese government bond yield and the total assets of the Bank of Japan. In addition, we also

[1] Source: Japan Financial Services Agency.

[2] Bankrupt loans are loans to borrowers in legal bankruptcy and past due loans by 6 months or more. Restructured loans are named after the sum of past due loans by 3 months but less than 6 months and restructured loans. See Data section and Appendix A for more details.

control for bank competition by estimating the Boone indicator at bank level. Thirdly, to account for potential endogeneity, we employ the dynamic panel threshold analysis, where Generalised Methods of Moments type estimators are used (Kremer et al., 2013). This methodology allows us to examine whether these relationships are stable over the observed period (financial years from 2000 to 2014) which embraces quite a few important events. They are: the final phase of the banking crisis (2000-2001), the restructuring period (2001-2003), the presence of quantitative easing (2001-2006 and from 2010), the global financial crisis (2007-2008), and the Tohoku earthquake (2011). The advantage of dynamic threshold analysis is that structural events are not set exogenously, but are rather revealed by the threshold analysis. Lastly, we extend our analysis by using a panel vector autoregression (p-VAR) approach to address the underlying causality between quantitative easing and bank risk.

Our results show that quantitative easing reduces bank risk in terms of bankrupt and restructured loan ratios. However, quantitative easing appears to undermine overall bank stability. The results could entail the countervailing effects of quantitative easing on bank risk-taking (Buch et al., 2014; De Nicolò et al., 2010). Although quantitative easing mitigates the interest rate burden for borrowers and inflates the value of pledged assets, it may induce banks to engage in riskier projects in the search for higher yield. Along these lines, Jiménez et al. (2014) argue that low interest rates reduce the probability of default for current variable rate loans, but at the same time banks tend to issue new loans to borrowers with worse credit ratings. Regarding the causality between the variables of interest, the panel VAR analysis suggests that quantitative easing does indeed generate this causal relationship between risk and competition.

The paper proceeds as follows. Section 2 briefly reviews the literature and associated hypotheses. Section 3 presents the methodologies. Section 4 introduces the data. Section 5 discusses the results. Finally, section 6 concludes.

2. Related literature and the quantitative easing and risk hypothesis

In this section, we establish our research hypothesis based on the literature regarding the relationship between quantitative easing and bank risk. After a long history of nearly zero policy rates during the 1990s to avoid a deflationary slump (Leigh, 2010), the Bank of Japan initiated the policy of quantitative easing in March 2001 through long-term government bond purchases. Thereafter, assets purchased were broadened to private assets held by private banks, asset-backed securities and asset-backed commercial papers (Girardin & Moussa, 2011). Officially ended in March 2006, the first quantitative easing period did not firmly prove its effectiveness in extracting the economy from the deflationary cycle (Bowman et al., 2015; Ueda, 2012; Ugai, 2007).

The importance of quantitative easing has been addressed in relation to its potentially significant impact on aggregate demand, financial markets and economic growth (Bowman et al., 2015; Glick & Leduc, 2012; Schenkelberg & Watzka, 2013). Regarding its effect on the banking system, the bank lending channel is emphasised as a main conduit (Bowman et al., 2015; Hosono, 2006). As Lucas (2014) points out, the success of quantitative easing (in the US) is partly indicated by increased risk-taking and, hence, more bank lending. Starting with the zero lower bound interest rate policy, Hosono (2006) investigates the different impacts of expansionary monetary policy on bank lending. This paper addresses the three important bank characteristics, namely size, liquidity and capitalisation, which could alter a bank's reaction to the monetary policy stance. Results indicate that expansionary monetary policy in Japan is less effective for undercapitalised banks. Lending of small, less liquid and well-capitalised banks are more exposed to monetary policy shocks than their counterparts.

Inspired by Hosono (2006) but slightly more comprehensive is Bowman et al. (2015), which particularly focuses on the first quantitative easing period. Bowman et al. (2015) show that bank lending, consequent on quantitative easing, is transmitted through the liquidity channel. However, the results suggest that the liquidity injection of the central bank was inhibited by interbank illiquidity, thus the size of the boost to credit was relatively small. Unlike the findings of Hosono (2006), less-capitalised banks benefit more from quantitative easing than their well-capitalised peers. Weaker banks, in terms of higher nonperforming loan to asset ratios, also appear to be more sensitive to the liquidity injection. Bank size is reported to be insignificant in affecting the relationship between bank lending growth and liquidity. Kobayashi et al. (2006) also find evidence to support the conclusion that financially weaker banks and firms reap more benefits from quantitative easing through positive excess stock returns.

To this end, to the best of our knowledge, no study has established a clear link between quantitative easing in Japan and bank risk-taking using bank level information. Academics and policymakers have addressed the potentially disproportionate bank risk-taking associated with the enactment of quantitative easing (Chodorow-Reich, 2014; Claeys & Darvas, 2015). Quantitative easing is supposed to encourage financial institutions to attempt socially desirable risk-taking. However, banks may deviate from their secured path when excessive risk-taking is recorded (Claeys & Darvas, 2015). In addition, under lax lending standards and low interest rates, the likelihood is that the number of more risky borrowers being offered new loans could rise, and so could credit risk (Ioannidou et al., 2015; Jiménez et al., 2014). The countervailing effect of interest rate changes on bank risk is also addressed in Buch et al. (2014). Lower interest rates could reduce the cost burden for borrowers, increase the collateral value, and subsequently raise the likelihood of repayment. At the same time, borrowing capacity rises with higher collateral prices, and banks are induced to engage in riskier projects to offset their lower profits due to lower interest rates. On the contrary, Lucas (2014) argues that quantitative easing could unintentionally reduce bank risk-taking incentives. Banks benefit from the term premium in the yield curve if the maturity of their assets exceeds that of their liabilities. When the yield curve is flat, they may be discouraged to issue long-term loans which may be more desirable for borrowers.

In this regard, we leave our quantitative easing-risk hypothesis open: *The implementation of quantitative easing could lead to either an increase or a reduction in bank risk.*

3. Methodology - Dynamic panel threshold analysis

To examine the quantitative easing-risk nexus, we adopt the dynamic panel threshold model introduced by Kremer et al. (2013). This methodology allows for the estimation of a threshold effect within a panel data framework involving endogenous regressors. Apart from tackling endogeneity concerns, another advantage of this methodology in the case of Japanese banking is that no a priori assumption is needed with regard to structural breaks. Such breaks, within the threshold model, are endogenously determined from the underlying data generating process. The model estimates threshold values for quantitative easing over time, which in turn signify regime changes.

In some detail, the model specification is written as:

$$y_{it} = \mu_i + \beta_1 q_{it} I(q_{it} \le \gamma) + \delta_1 I(q_{it} \le \gamma) + \beta_2 q_{it} I(q_{it} > \gamma) + \varphi z_{it} + \varepsilon_{it} \qquad i = (1, \dots, N), t = (1, \dots, T)$$

where μ_i indicates bank-specific fixed effect[3]; $I(.)$ is the indicator function indicating the regime defined by the threshold variable (q_{it}) and the threshold level γ; q_{it} is both the threshold variable and the regime-dependent regressor, whereas z_{it} is a vector of control variables, which may include both endogenous and exogenous variable.

As in Kremer et al. (2013), we account for the regime intercept (δ_1) because omitting the intercept may result in biases in the threshold estimates and regression slopes (Bick, 2010). ε_{it} is the error term.[4] As in Caner and Hansen (2004) and Kremer et al. (2013), we estimate equation (1) using GMM to account for endogeneity. The first lag of the endogenous variable is used as the instrument.

4. Data

Our data are extracted from semi-annual financial reports of Japanese commercial banks published on the Japanese Bankers Association website. Our sample consists of 3491 observations for financial years 2000 to 2014. Three particular types of commercial banks are examined, namely City Banks, Regional Banks I and Regional Banks II. They form more than half the banking system and correspond to various types of operations. If City Banks are more involved in different aspects of banking business, Regional Banks are prone to

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[3] To eliminate bank-specific fixed effects, as suggested by Kremer et al (2013), we employ the forward orthogonal deviations transformation proposed by Arellano and Bover (1995).

[4] The estimation procedure is as follows. First, a reduced-form regression is estimated for endogenous variables as a function of the instrumental variables. Second, using least squares, we estimate equation (1) for a fixed threshold with the predicted values of endogenous variables obtained from the first step regression. Third, the second step regression is repeated to find the estimator of the threshold value associated with the smallest sum of squared residuals. The critical values for the 95% confidence intervals of the threshold value are: $\Gamma = \{\gamma : LR(\gamma) \le C(\alpha)\}$, with $C(\alpha)$ is the 95% percentile of the asymptotic distribution of the likelihood ratio statistic $LR(\gamma)$ (Caner & Hansen, 2004). The slope coefficients are estimated by GMM procedure for the formerly used instruments and estimated threshold.

conventional banking activities. City Banks are referred as *Main Banks* in the *horizontal keiretsu* network – the enterprise groups consisting of one large firm for every major sector pre- and post-crisis. These banks act as the core of the business group and offer venture capital for affiliates. The number of City Banks has declined over time since the crisis occurred in the 1990s. Besides, during the restructuring period, City Banks benefited from mergers which empowered their resistance to overcome the consequences of the crisis[5].

The operating locations of Regional Banks are refined by their scope of business, with smaller geographic region restriction for Regional Banks II. These banks are the smallest in comparison with the other two. Unlike City Banks, Regional Banks mainly invest in government bonds and originate loans for small and medium firms in their specific areas where their head offices are located. Thus, Regional Banks are more committed to the local development of the prefectures. There are other different kinds of banks currently operating in Japan, for example, Trust Banks, Long-Term Credit Banks, *Shinkin* banks (credit cooperatives), and foreign banks. Due to data unavailability or differences in business features, we do not include non-commercial banks in our study.

As dependent variables representing bank risk-taking, we opt for bankrupt loans to total assets (BRL ratio), restructured loans to total assets (RSL ratio), and the natural logarithm of Z-score[6]. The first two variables characterise credit risk, whereas the remaining variable is a proxy for overall bank stability. They are incorporated respectively in the model to analyse the highlighted hypotheses. Bankrupt and restructured loans are obtained from data of risk-monitored loans disclosed under the Banking Law (see Appendix A). Bankrupt loans are the sum of bankrupt loans and non-accrual loans,[7] while restructured loans are the sum of past due loans by 3 months or more but less than 6 months, and restructured loans.[8] The ratios of these risk-monitored loans to assets capture credit risk, similar to the

[5] Mitsui Bank and Taiyo-Kobe Bank to form Sakura Bank; Fuji, Dai-Ichi Kanyo, and Industrial Bank of Japan to form Mizuho Bank; Sanwa and Tokai Banks to form UFJ Banks; UFJ Banks and Bank of Tokyo-Mitsubishi; Sumitomo Bank and Sakura Bank (Nakamura, 2006) .

[6] Nonperforming loan ratio and Z-score are used extensively in the literature to represent bank risk (Agoraki et al., 2011; Beck, 2008; Buch et al., 2012).

[7] Reported in Japanese commercial banks' balance sheets, these loans are named loans to borrowers in legal bankruptcy, and past due loans in arrears by six months or more.

[8] The Japanese Bankers Association originally defined restructured loans as loans of which interest rates were lowered. In 1997, the definition was extended to loans with any amended contract conditions and loans to corporations under on-going reorganisation (Montgomery & Shimizutani, 2009).

nonperforming loan to asset ratio that has been widely used in the literature to test for the *competition-fragility* nexus (Beck, 2008). Bank stability indicated by the Z-score is another gauge for the likelihood of bank failure (Beck et al., 2013; Laeven & Levine, 2009). This is defined as the number of standard deviations below the mean of return on assets that would result in insolvency by evaporating capital ($Z_{it} = \left(ROA_{it} + Capital\ ratio_{it}\right) / \sigma ROA_{it}$) (Boyd & De Nicolo, 2005).

To ensure the robustness of our estimation, we analyse the dynamic panel threshold model with two proxies for competition, the Boone indicator and the Lerner index, with the former being our primary interest. With regard to quantitative easing, we choose the bank-specific lending rate calculated as interest income on loans divided by loans and bills discounted[9]. We employ the bank-specific lending rate as the threshold variable for quantitative easing for several reasons. First, under the zero lower bound, short-term interest rates are inoperative (Girardin & Moussa, 2011). Second, the Bank of Japan loan rate, uncollateralised overnight call rates, and the Bank of Japan's total reserves, the amount of asset purchases, and government bond yields do not reflect individual bank characteristics in relation to changes in quantitative easing. Third, we could avoid aggregation bias and enhance the compatibility of quantitative easing proxy with the dependent variable and the Boone indicator. We also control for the impact of competition. As quantitative easing influences deposit interest rates, it may in turn affect bank competition in the loan market. In addition, one may argue that lending rate is not a direct measure of quantitative easing, and may be affected by competition. To tackle the potential endogeneity between the three main variables of interest, we treat them as endogenous in the dynamic panel threshold model.[10] For robustness, we also use the 10-year Japanese government bond yield and the Bank of Japan total assets (Lyonnet & Werner, 2012) as alternative proxies for quantitative easing.

Regarding a subset of explanatory variables, we specify a number of control variables varying from bank characteristics to macroeconomic impact. To account for capitalisation

[9] We could also use the amount of asset purchases or Japanese government bond's yield as measures for quantitative easing (Bowman et al., 2015; Lyonnet & Werner, 2012; Voutsinas & Werner, 2011).

[10] We use the first lag of the endogenous variable as its instrument to preserve information. Following Kremer et al. (2013), all available lags of the endogenous regressor are also examined. In fact, the corresponding results reveal little variation in the parameters estimated.

and the potential moral hazard problem, we use the capital to assets ratio (Tabak et al., 2012)[11]. Bank size is taken as the natural logarithm of total assets. We also take into consideration the impact of revenue diversification which is the ratio of non-interest income to total operating income (Anginer et al., 2014; Beck et al., 2013), assets diversification represented by the ratio of securities to assets (Zhang et al., 2013), and liquidity which is defined as liquid assets[12] to total assets (Jeon et al., 2011). GDP growth is included to reflect the influence of macroeconomic environment (Jiménez et al., 2013). Market capitalisation accounts for financial market development and also functions as an alternative source of funds for incumbent firms (Beck et al., 2013). Descriptive statistics of data used are displayed in Table 1.

[INSERT TABLE 1 ABOUT HERE]

We also measure competition at bank level using the Boone indicator (see Appendix B for the exposition of the Boone indicator). The average Boone indicator for the entire sample from 2000 to 2014 is estimated at -0.0542.[13] Reported in Mirzaei and Moore (2014), the average Boone indicator for Japanese banking between 1999 and 2011 is -0.02.[14] Figure 1 illustrates the mean value of the Boone indicator over time for all banks in our sample. Its highest score in absolute value is recorded in March 2002 at -0.0813, indicating the toughest degree of competition for the entire period. During the restructuring period (September 2000 to March 2003), the government imposed policy changes on the banking system in order to revitalise its resilience to the aftermaths of the crisis. In addition, undercapitalisation and the threat of nonperforming loans induced fragile banks to agree to merger proposals from financially healthier banks. The consolidation tendency was augmented by a number of mergers between large City Banks, indicating an adverse phase for *too-big-to-fail* banks in maintaining their market power. Afterwards, the average score slightly increased to -0.0483 in March 2004, only to remain relatively constant until

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[11] As capital ratio is part of the formula of Z-score, we exclude it from models in which *ln*Z-score is used.

[12] Liquid assets = Cash and due from banks + call loans + receivables under resale agreements + receivables under securities borrowing transactions + bills bought + monetary claims bought + trading assets + trading account securities + money held in trust (Radić, 2014).

[13] Delis (2012) includes Japan in the sample of 84 countries. The average Boone indicator for Japanese banks during 1988-2005 is -0.584.

[14] Note that the data are obtained from World Bank for the whole banking system.

September 2008. This may serve as evidence in supporting the positive outcomes of government intervention. Within that time frame, the turbulence caused by the huge amount of nonperforming loans had been alleviated gradually.

[INSERT FIGURE 1 ABOUT HERE]

There was a shift in the Boone indicator which signified higher competition during the US subprime crisis. At the end of March 2009, the corresponding Boone indicator dropped from -0.0489 (in September 2008) to -0.0659. The contagion of the global financial crisis possibly caused Japanese banks to deviate from their profit goals. The deterioration in profit, in turn, could reduce bank market power. Between September 2009 and September 2012, the Boone indicator had a similar stable trend as after the restructuring period, before slightly decreasing towards the end of the sample period.

The corresponding stiff competition is identified by higher absolute values of the Boone indicator. There is no specific benchmark for the value of β in general, yet what we have found implies a moderate degree of competition in the banking sector, as the figures are not too distant from zero. Table 2 provides further insight of competition among each bank type. In general, competition within City Banks (-0.0654) and Regional Banks II (-0.0559) are more intense than between Regional Banks I (-0.0518). The largest magnitude (absolute value) of Boone indicator is recorded for City Banks in March 2002 at -0.1906. The trend of competition in City Banks during the restructuring period is more volatile than those in the other two types, indicating the effect of the aforementioned consolidation tendency. The onset of the US credit crunch 2007-2008 seemed to impose a pronounced effect on competition between Regional Banks II, notably at -0.0809 in March 2009. A potential explanation could rest on the size factor which may denote a bank's resistance to external shocks. Regional Banks II are the smallest compared to the other two and operate under more limited geographic restrictions. Hence, the potentially high likelihood that Regional Banks II being more exposed to exogenous shocks would erode their profits and weaken their market power. Nevertheless, competition in Regional Banks II appeared to be more relatively stable compared to City Banks and Regional Banks I after the global financial crisis.

We also compute the Lerner index for comparison purposes (see Appendix B for its estimation). The trend of the Lerner index over time illustrated in Figure 1 concurs with the previous findings of the Boone indicator.[15] Our result reveals that the average Lerner index is 0.2565, with some variation across bank types (the average Lerner index reported for Japanese banks from 2003 to 2010 in Fu et al. (2014) is 0.2521). The level of competition is relatively tougher for City Banks (0.1467) and Regional Banks II (0.2421) than for Regional Banks I (0.2777), in line with the rank of Boone indicators formerly reported for three types. The trend of the Lerner index over time is very similar to the pattern of the Boone indicator (see Fig. 1). The two points expressing the strongest competitive environment are also observed in March 2002 and March 2009. Our results, however, are different from findings of Liu and Wilson (2013), possibly because they obtain the Lerner index by estimating the whole banking system, including Trust Banks, *Shinkin* Banks and Credit cooperatives during 2000-2009. Concerning the three types in our sample, Liu and Wilson (2013) find that City Banks have the greatest market power, followed by Regional Banks I and Regional Banks II. A similar interpretation is drawn by Montgomery et al. (2014) as large banks enjoy greater market power post-mergers.

[INSERT TABLE 2 ABOUT HERE]

5. Results

5.1 Quantitative easing and risk: quantitative easing as the threshold variable

Findings for the risk-quantitative easing nexus are reported in Table 3. There exists a positive relationship between lending rate and risk in all different model specifications (columns 1 to 3). A rise in the lending rate is found to increase bankrupt/restructured loan ratios and *ln*Z-score, statistically significant in both regimes. When risk is measured by the bankrupt loan ratio and the restructured loan ratio, the threshold value is identified at 1.2052% (column 1) and 1.0562% (column 2). When the *ln*Z-score is used, the threshold value is 0.9401% (column 3). To this end, quantitative easing is beneficial in terms of reducing credit risk. The effect is more prominent in the high regimes, where the

[15] There are some exceptional cases when market power characterised by the Lerner index is negative, but occasionally found. Agoraki et al. (2011) and Fu et al. (2014) explain the implication of negative Lerner index by the non-optimising behaviour of banks which are unable to price above marginal cost.

coefficients are 0.037 (column 1) and 0.0385 (column 2). Although the coefficients indicating the impact of the lending rate on risk-monitored loan ratios in the low regimes are statistically significant, the magnitude is quite negligible (0.0088 and 0.0063 in columns 1 and 2, respectively). Nevertheless, this favourable effect of lower risky loan ratios associated with quantitative easing may come at the expense of bank stability. The reason is that, as reported in column 3, the *ln*Z-score is also reduced, given an aggressive quantitative easing policy. Notably, the magnitude of the coefficients for the impact of the lending rate provides insightful implications. Compared to the detrimental effect that quantitative easing could impose on bank stability (0.3416 and 0.2833 in the low and high regimes, respectively), the beneficial impact that it exerts on credit risk is quite small. In a nutshell, comparing the results to the hypotheses set out in section 2, we can conclude that quantitative easing could lower credit risk but may harm overall bank stability.

In an attempt to explain the aforementioned findings, the implication of variability in our results could be interpreted by the countervailing effect of low interest rates on bank risk-taking as discussed in Buch et al. (2014). On the one hand, quantitative easing may reduce risk, as it aims to facilitate lending so that increased investment can boost economic growth. Both banks and borrowers can benefit from ample liquidity injected by quantitative easing to strengthen their resistance to exogenous shocks. Low interest rates would encourage more potential borrowers to apply for funding because of a greater probability of fulfilling their repayment duties. As evidenced in Jiménez et al. (2014), a low interest rate reduces the interest burden on existing loans for borrowers. Therefore, lower bankrupt/restructured loan ratios would be expected. On the other hand, quantitative easing could amplify risk. When banks foresee an extended period of low interest rate, they may alter their risk-taking appetite towards riskier projects to pursue greater gains (Altunbas et al., 2014; Gambacorta, 2009). In more detail, low yield and abundant liquidity accelerate asset prices and promote leverage, in turn inducing excessive risk-taking (Dell'Ariccia et al., 2010). A larger loanable proportion of collateral and the search for yield (Rajan, 2005) may drive banks to grant more risky loan portfolios (Jiménez et al., 2014), or to invest in higher yield-higher risk instruments. Another risk-taking channel could be through a typical type of moral hazard, where banks realise the continuity of quantitative easing policy in difficult economic times. As Altunbas et al. (2014) argue, banks may perceive the presence of a so-called *insurance effect*, in which monetary easing during a financial downturn is expected to decelerate the fall of asset values. The prediction of a lower probability of large downside risk, therefore, would magnify bank risk-taking. This perception may well be the case of prolonged low interest rate and extensive quantitative easing in Japan. Taken together, these arguments could explain for lower bank stability corresponding to quantitative easing.

[INSERT TABLE 3 ABOUT HERE]

Interestingly, in terms of control variables, the results reveal a negative association between the Boone indicator and risk-monitored loan ratios. Hence, greater competition would be harmful for banks because of higher bankrupt/restructured loan ratios. This finding is reinforced by the positive association between the Boone indicator and the *ln*Z-score in column 3, indicating higher bank stability under conditions of lower competition. In particular, in terms of diminishing risk and enhancing bank soundness, there are four variables: the capital ratio, asset diversification, liquidity and market capitalisation. In contrast, higher GDP growth is found to increase risk-monitored loan ratios, probably due to the softened lending standards during good economic times (Dell'Ariccia & Marquez, 2006). Turning to the Z-score, there is a favourable impact on bank stability during economic upturn and when banks divert their focus to noninterest income. Bank stability increases with larger bank size, while the bankrupt loan ratio decreases.

The number of banks in each regime is shown in Table 4. Analysing the trend of the number of observations in column 1, we observe a significant increase of banks in the low regime after the global financial crisis. Especially, from March 2011 to March 2015, almost all banks in the sample charged less than 1.2052% lending rate. Note that this time frame covers the on-going quantitative easing policy (since October 2010). Illustrated in column 3, the distribution of the number of banks in the low regime provides further evidence for the initial quantitative easing period. Recall that the threshold value for column 3 is 0.9401%, which is lower than the values for columns 1 (1.2052%) and 2 (1.0562%). From September 2003 to March 2006, the number of banks charging lending rate lower than 0.9401% increased monotonically. This tendency indicates the effect of the first quantitative easing period (March 2001-March 2006). In the high regimes of all model specifications, it is confirmed that the number of observations gradually decreased during this period.

[INSERT TABLE 4 ABOUT HERE]

In Table 5, we use the 10-year Japanese government bond yield (columns 1 to 3) and Bank of Japan assets (columns 4 to 6) to replace lending rate as the threshold variable. The results show a positive influence of bond yield on the *ln*Z-score in both regimes of column 3. This is in line with the previous findings that quantitative easing reduces bank stability (reported in column 3 of Table 3). The magnitude of the impact in the high regime (0.6513, column 3) is also notable compared to other models. Interestingly, the bond yield affects risk-monitored loan ratios differently in two regimes. There is a statistically significant positive relationship between the bond yield and the bankrupt/restructured loan ratios in the high regimes. The coefficients on the bond yield's impact are 0.0131 (column 1) and 0.0107 (column 2). This relationship turns out negative in the low regimes (-0.017 in column 1 and -0.0149 in column 2). It is also worth noting that the absolute magnitudes of the impact of bond yield on risk-monitored loans in the two regimes are approximately the same (around 0.01). Additionally, the threshold value is consistently realised at 1.032%. Thus, when the bond yield is below 1.032%, quantitative easing increases credit risk. In this regard, more aggressive quantitative easing would encourage banks to take on more risk. First, banks may tend to soften lending standards due to low yield and interest rate, thereby issuing loans to less creditworthy borrowers (Jiménez et al., 2014). Second, as Ioannidou et al. (2015) argue, due to low monetary policy rate, banks may be less concerned about the compensation which should be required for the higher risk taken. In fact, Ioannidou et al. (2015) find that during monetary expansion, banks charge riskier borrowers relatively less than what they should. When the bond yield is greater than 1.032%, quantitative easing reduces credit risk, similar to our previous conclusion drawn from lending rate (Table 3).

[INSERT TABLE 5 ABOUT HERE]

The bond yield is lower than the threshold value (1.032%) in March 2003, September 2010, and from September 2011 to March 2015. The last time frame includes the current quantitative easing period. If we combine this finding with the aforementioned impact of the threshold, the on-going quantitative easing may pose a threat to the banking system by augmenting credit risk. Regarding control variables, similar to the results reported in Table 3, we also find that competition increases credit risk and bank fragility. Larger size, a higher capital ratio, more liquidity, greater asset diversification, revenue diversification and market capitalisation would help lower credit risk. Higher GDP growth, on the other hand, would increase credit risk exposure. In terms of bank stability, it would be enhanced with larger bank size, more diversified income, higher GDP growth, and greater market capitalisation.

A first glance at columns 4 to 6, where the Bank of Japan total assets are used as a proxy for quantitative easing, reveals a consistent estimate of the threshold value at JPN 118,437,502 million. There is a negative association between the Bank of Japan's balance sheet size and risk variables. For credit risk, this relationship is statistically significant in the low regime (-0.0347 in column 4 and -0.0286 in column 5), implying a favourable impact of quantitative easing. The influence of the Bank of Japan's balance sheet size on risk-monitored loan ratios in the high regime is insignificant. Differently, for bank stability, when the Bank of Japan's balance sheet is greater than the threshold, a more aggressive quantitative easing policy would reduce bank soundness (-0.1709 in column 6). The relationship between quantitative easing and bank stability is insignificant in the low regime. These results strengthen those reported in columns 1 to 3, where the bond yield is the proxy for quantitative easing. Up to a certain level of asset purchases (JPN 118,437,502 million), quantitative easing lessens credit risk. When the amount of asset purchases passed the threshold, quantitative easing reduces bank stability.

The time frame in each regime complements these findings. First, the periods of high regimes coincide with the two quantitative easing periods. In particular, the amount of asset purchases, which were higher than the threshold, are recorded from March 2001 to March 2006, and from March 2011 to March 2015. Hence, the additional asset purchases of the Bank were not really effective due to its detrimental impact on bank stability. Second, the period of low regimes falls in to the gap between the two quantitative easing periods, and also embraces the global financial crisis. During this interval (September 2006-September 2010), more asset purchases would have mitigated credit risk. However, overall, the estimated impact suggests that the reduction in credit risk may not be considerable compared to the reduction in bank soundness (e.g. -0.0347 in column 4 versus -0.1709 in column 6). The influence of other control variables appears consistent as previously reported in columns 1 to 3 and in Table 3.

5.2 The panel VAR specification

Given some variability in our results, which could be driven by endogeneity issues, we attempt to address the underlying dynamics between risk and quantitative easing. We adopt the methodology of panel vector autoregression (p-VAR) to account for the causality relationship as well as the existence of unobservable heterogeneity, specified by an individual specific term. An advantage of the model is its lack of assumptions about the relationship between variables. We treat all three variables in the equation system as endogenous.[16] Risk, proxied by the bankrupt loan ratio, the restructured loan ratio, and the *ln*Z-score, is incorporated respectively in the analysis. The Boone indicator and lending rate are proxies for competition and quantitative easing, respectively. We also include bank size as an exogenous control variable because of its importance in the Japanese banking structure. As discussed in the Data section, City Banks are the biggest and operate in a wide range of geographic regions, whereas Regional Banks II are the smallest. The nature of banking business also varies across three types. Besides, *two-big-to-fail* City banks are at the centre of the *keiretsu* network as well as being the important nodes channelling the impact of quantitative easing.

Following the estimation of panel VAR, we derive the Impulse Response Functions (IRFs) (Fig. 2 to 4), which enable us to interpret the reaction of one variable to a shock in another variable in the system. We also report the Variance Decomposition (VDCs) for forecast horizons of 5 and 10 periods to illustrate the variance of the response variable to a shock in another variable (Table 6). All model specifications satisfy stability condition.[17]

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[16] Following Love and Ariss (2014), we run the model on lag order 1 to preserve information.

[17] The variables enter the equation system as endogenous, with the most exogenous ones appearing first (Love & Zicchino, 2006). Following Love and Zicchino (2006), fixed effects are removed by using the *Helmert procedure* (Arellano & Bover, 1995).

The first order VAR model takes the form of: $w_{it} = \mu_i + \Phi w_{it-1} + e_{i,t} \quad i = 1,...,N; \qquad t = 1,...,T$

where w_{it} is a vector of three random variables: quantitative easing *QE*, Competition *Comp* and risk *R* (bankrupt loan ratio, restructured loan ratio and *ln*Z-score), Φ is a 3x3 matrix of coefficients, μ_i is a vector of *m* individual effects, and $e_{i,t}$ is a multivariate white-noise vector of *m* residuals. The equation system to be estimated with lag order one is as follows:

$$
\begin{aligned}
QE_{it} &= \mu_{10} + a_{11}QE_{it-1} + a_{12}Comp_{it-1} + a_{13}R_{it-1} + e_{1i,t} \\
Comp_{it} &= \mu_{20} + a_{21}QE_{it-1} + a_{22}Comp_{it-1} + a_{23}R_{it-1} + e_{2i,t} \\
R_{it} &= \mu_{30} + a_{31}QE_{it-1} + a_{32}Comp_{it-1} + a_{33}R_{it-1} + e_{3i,t}
\end{aligned}
$$

[INSERT FIGURE 2 ABOUT HERE]

[INSERT FIGURE 3 ABOUT HERE]

In terms of the risk-quantitative easing nexus, in the short term, there is a positive and significant response of risk-monitored loan ratios to a one standard deviation shock in lending rate (Fig. 2-3, last row, first column). This positive reaction is similar to the findings in Table 3. There is no evidence for a significant response of bank stability to a shock in the lending rate. The diagrams also reveal insignificant responses of lending rate to shocks in risk variables. Thus, if there exists a positive shock in lending rate, which translates into decreased quantitative easing, credit risk could rise accordingly. Hence, the simulation base of panel VAR could reinforce the claim of lower credit risk as a result of quantitative easing.

[INSERT FIGURE 4 ABOUT HERE]

Regarding the risk-competition nexus, a shock to the Boone indicator has a negative and significant impact on bankrupt loan ratio (Fig.2, second row, first column). In terms of reverse causality, shocks in risk variables generate insignificant responses of the Boone indicator. Two scenarios can be at play to interpret the results. A positive shock in the Boone indicator which denote lower competition will lead to decreased credit risk and increased bank stability. Turning to the competition-quantitative easing linkage, a shock in the lending rate would generate a negative response in the Boone indicator, marginally significant in the short-run (Fig. 2-4, last row, second column).

Complementing the findings of the IRFs, the VDCs show that changes in competition are important in explaining the variation in the bankrupt loan ratio (5.66%), the restructured loan ratio (1.26%) and the *ln*Z-score (12.79%) (Table 6, 10 periods). In contrast, about 0.36% and 12.4% variation in the Boone indicator is due to innovations from the restructured loan ratio and the *ln*Z-score, respectively. Variation in the bankrupt loan ratio does not explain the variation in competition at all. Findings from the IRFs and VDCs reveal that competition triggers its relationship with risk.

[INSERT TABLE 6 ABOUT HERE]

Regarding the risk-quantitative easing relationship, about 46.79% variation in bankrupt loan ratio is explained by variation in quantitative easing, while only 0.03%

variation in quantitative easing is explained by shocks in bankrupt loan ratio. Similarly, 30.89% variation in restructured loan ratio is due to shocks in quantitative easing, while 0.07% variation in quantitative easing is explained by changes in restructured loan ratio. Differently, changes in quantitative easing is not so important in explaining the variation of bank stability. The reason is that while 3.01% variation in lending rate is due to innovations in *ln*Z-score, only 0.68% variation in *ln*Z-score is attributed to variation in lending rate. To this end, along with results from the IRFs, quantitative easing is found to originate its relationship with risk.

The variation in the Boone indicator indicated by variation in lending rate is distinguishably larger than the variation in lending rate explained by changes in the Boone indicator (8.35%, 6.14%, 10.3% in comparison to 0.01%, 0.01%, 0.12% in columns 1-3, respectively). A conclusion of the causality starting from quantitative easing to competition can be drawn.

6. Conclusion

Our findings indicate that in an environment where quantitative easing is taking place, banks might find it more challenging to compete with their counterparts. To improve their competitiveness, banks could strengthen their competence from other aspects, e.g. capitalisation, liquidity, and asset diversification. Bank executives could enhance banking services by, e.g., diversifying their investments or increasing unconventional business activities to offer more benefits to their customers in time and cost savings. In addition, focusing on relationship banking, improving their flexibility in debt rollover, and operating more efficiently may also be among the tactics bringing banks ahead their rivals. The proposed threshold values for lending rates in this study may also be useful for bank managers to construct their risk management policy.

For policymakers, e.g. the Japan Financial Services Agency, relaxing entry and exit for the banking industry, promoting small and medium sized banks, or disentangling business operation restrictions could create a competitive environment which in turn would diminish bank risk-taking. Policymakers could also encourage the mutual assistance prevailing under the *keiretsu* network. Note that a disadvantage of *keiretsu* affiliation is that main banks could exert their monopoly power in loan financing. Our analyses show that attempts to discourage competition increase credit risk. Therefore, our results argue that *keiretsu* should be applied with extreme caution.

Last but not least, to take into account the stability of the banking system while exercising quantitative easing, regulators may revise rules associated with the initial credit screening and barriers in lending principles. In more details, avoiding incorrect evaluations at the beginning of the loan generating process and complying with lending standards help banks lessen the possibilities of future uncertainty. These policies should not contradict but promote the efficacy of quantitative easing and *Abenomics* - the current monetary and economic growth policy. Given that the Bank of Japan has adopted negative interest rate in January 2016 for the first time in its history, Japan would warrant a very interesting platform for future research. If the negative interest rate could drive economic recovery, it would open up a new era for monetary policy.

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Table 1. Descriptive Statistics

Variable	Mean	S.D.	Min	Max
BRL ratio	0.0263	0.0219	0.0000	0.6765
RSL ratio	0.0092	0.0093	0.0000	0.1958
*Ln*Zscore	3.9335	0.5223	0.0000	5.6410
Boone indicator	-0.0542	0.0579	-1.6390	-0.0391
Lerner index	0.2565	0.3365	-4.0314	0.7583
Lending rate	0.0106	0.0024	0.0012	0.0366
Size	14.5717	1.1591	12.0571	19.0109
Capital ratio	0.0432	0.0240	-0.7882	0.1279
Asset diversification	0.2394	0.0770	0.0000	0.4807
Liquidity ratio	0.0722	0.0380	0.0089	0.3679
Revenue diversification	0.2220	0.0817	0.0577	0.5445
GDP growth	0.0032	0.0234	-0.0787	0.0543
Bond yield	0.0122	0.0039	0.0041	0.0185
Bank of Japan assets	18.7201	0.2530	18.3119	19.5192
Market capitalisation	19.2235	0.2546	18.8248	19.6968

Notes: This Table reports the descriptive statistics for key variables employed in the dynamic panel threshold analysis. Number of observations: 3491. BRL ratio: bankrupt loans to assets, RSL ratio: restructured loans to assets, Z-score $Z_{it} = \left(ROA_{it} + Capital\ ratio_{it}\right) / \sigma ROA_{it}$, lending rate=interest income on loans/loans and bills discounted, size=*ln*(total assets), capital ratio=equity/assets, asset diversification=securities/assets, liquidity=liquid assets/total assets, revenue diversification=non-interest incomes/operating income, Bond yield: 10-year Japanese government bond yield, Bank of Japan assets and market capitalisation is in natural logarithm. S.D.: Standard deviation.,

Table 2. Boone Indicator and Lerner Index per Bank Type.

Variable	Boone			Lerner		
Time	City	Regional 1	Regional 2	City	Regional 1	Regional 2
Sep-00	-0.0840	-0.0575	-0.0581	0.1337	0.2001	0.1732
Mar-01	-0.1129	-0.0597	-0.0683	-0.1054	0.1321	0.1356
Sep-01	-0.1332	-0.0605	-0.0886	-0.2174	0.1146	0.0830
Mar-02	-0.1906	-0.0625	-0.0894	-0.7096	0.0647	0.0289
Sep-02	-0.0517	-0.0535	-0.0795	0.0709	0.1813	0.1597
Mar-03	-0.1206	-0.0515	-0.0810	-0.8319	0.1173	0.0257
Sep-03	-0.0429	-0.0699	-0.0469	0.2166	0.2190	0.2822
Mar-04	-0.0616	-0.0452	-0.0503	0.1374	0.3101	0.2645
Sep-04	-0.0601	-0.0469	-0.0455	0.1992	0.3177	0.3030
Mar-05	-0.0562	-0.0461	-0.0510	0.0917	0.3196	0.2993
Sep-05	-0.0432	-0.0486	-0.0458	0.3346	0.3169	0.3385
Mar-06	-0.0432	-0.0457	-0.0563	0.3082	0.3628	0.3086
Sep-06	-0.0413	-0.0464	-0.0550	0.3191	0.3303	0.2757
Mar-07	-0.0467	-0.0453	-0.0579	0.2326	0.3249	0.2593
Sep-07	-0.0546	-0.0496	-0.0468	0.2172	0.2880	0.2889
Mar-08	-0.0574	-0.0489	-0.0539	0.2117	0.2482	0.2143
Sep-08	-0.0513	-0.0453	-0.0537	0.0899	0.1893	0.1809
Mar-09	-0.0598	-0.0562	-0.0809	-0.0583	-0.0060	-0.0608
Sep-09	-0.0428	-0.0433	-0.0452	0.1706	0.2947	0.2577
Mar-10	-0.0445	-0.0443	-0.0493	0.2579	0.2994	0.2286
Sep-10	-0.0483	-0.0440	-0.0453	0.3296	0.3270	0.3011
Mar-11	-0.0435	-0.0442	-0.0442	0.2450	0.2942	0.2965
Sep-11	-0.0458	-0.0438	-0.0446	0.3260	0.3461	0.3194
Mar-12	-0.0579	-0.0464	-0.0438	0.3335	0.3280	0.3254
Sep-12	-0.0513	-0.0436	-0.0435	0.2787	0.3167	0.3245
Mar-13	-0.0586	-0.0605	-0.0426	0.3910	0.3630	0.3309
Sep-13	-0.0916	-0.0519	-0.0437	0.4174	0.3879	0.3878
Mar-14	-0.0419	-0.0612	-0.0439	0.3781	0.3673	0.3775
Sep-14	-0.0450	-0.0642	-0.0430	0.4285	0.4013	0.3927
Mar-15	-0.0438	-0.0663	-0.0451	0.3356	0.3895	0.3878
Total	-0.0654	-0.0518	-0.0559	0.1467	0.2777	0.2421

Notes: This Table reports the average Boone indicator and the Lerner index per bank type over time. Sep: September; Mar: March; 00-15: 2000-2015.

Table 3. Dynamic Panel Threshold Analysis for the Quantitative Easing-Risk Nexus (lending rate).

	1		2		3	
Dependent variable	BRL ratio		RSL ratio		lnZ-score	
Threshold estimates	1.2052%		1.0562%		0.9401%	
95% confidence interval	[0.9861% 1.2102%]		[0.9613% 1.1205%]		[0.9216% 1.0122%]	
Impact of threshold variables	Est.	S.e.	Est.	S.e.	Est.	S.e.
Low regime	0.0088***	0.0026	0.0063***	0.0020	0.3416***	0.0442
High regime	0.0370***	0.0090	0.0385***	0.0036	0.2833***	0.0751
Intercept	-0.1288***	0.0417	-0.1469***	0.0194	0.3061	0.2818
Impact of covariates						
Boone	-0.1359***	0.0557	-0.079***	0.0234	0.937***	0.3350
Size	-0.014***	0.0031	-0.0014	0.0022	0.1029***	0.0356
Capital ratio	-0.5453***	0.1079	-0.0560	0.0372		
Asset diversification	-0.0791***	0.0088	-0.0387***	0.0055	0.1920	0.1889
Liquidity	-0.0619***	0.0110	-0.0243***	0.0076	0.1413	0.2368
Revenue diversification	-0.0068	0.0059	-0.0061	0.0042	0.2777***	0.0771
GDP growth	0.0387***	0.0083	0.0297***	0.0050	0.2827***	0.0602
Market capitalisation	-0.002**	0.0009	-0.0003	0.0005	0.0374***	0.0107
Obs in low regime	2352		1789		1090	
Obs in high regime	959		1702		2401	

Notes: This Table reports the results from the dynamic threshold analysis using the first lag of the endogenous variable (Boone) as its instrument. The threshold variable is lending rate. BRL ratio: bankrupt loans to assets, RSL ratio: restructured loans to assets, Z-score $Z_{it} = \left(ROA_{it} + Capital\ ratio_{it}\right) / \sigma ROA_{it}$, lending rate=interest income on loans/loans and bills discounted, size=ln(total assets), capital ratio=equity/assets, asset diversification=securities/assets, liquidity=liquid assets/total assets, revenue diversification=non-interest incomes/operating income, market capitalisation is in natural logarithm. Est.: estimate, S.e.: standard error, Obs: number of observations. ***,**,*: significance at 1%, 5%, 10% level.

Table 4. Number of Observations in Each Regime for the Risk-Quantitative Easing Nexus.

	1		2		3	
	BRL ratio		RSL ratio		lnZ-score	
	1.2052%		1.0562%		0.9401%	
	Low	High	Low	High	Low	High
Sep-00	46	82	5	123	0	128
Mar-01	45	80	2	123	0	125
Sep-01	58	70	18	110	1	127
Mar-02	60	65	35	90	2	123
Sep-02	63	64	35	92	5	122
Mar-03	65	56	36	85	9	112
Sep-03	68	53	41	80	14	107
Mar-04	74	46	44	76	15	105
Sep-04	71	49	47	73	16	104
Mar-05	78	41	54	65	25	94
Sep-05	84	35	55	64	31	88
Mar-06	90	27	61	56	41	76
Sep-06	92	25	61	56	33	84
Mar-07	80	36	53	63	19	97
Sep-07	76	39	38	77	5	110
Mar-08	75	39	37	77	2	112
Sep-08	79	36	38	77	5	110
Mar-09	87	27	52	62	10	104
Sep-09	94	20	66	48	25	89
Mar-10	96	16	71	41	43	69
Sep-10	97	14	78	33	50	61
Mar-11	102	9	82	29	62	49
Sep-11	103	8	86	25	68	43
Mar-12	106	5	91	20	75	36
Sep-12	104	6	94	16	78	32
Mar-13	106	4	98	12	85	25
Sep-13	107	3	99	11	88	22
Mar-14	108	2	101	9	90	20
Sep-14	109	1	105	5	96	14
Mar-15	109	1	106	4	97	13
Obs	2532	959	1789	1702	1090	2401

Notes: This Table reports the number of observations in each regime over time for the risk-quantitative easing nexus, with lending rate being the threshold variable, and the Boone indicator being the proxy for competition. Threshold values of lending rate are obtained from the dynamic threshold analysis, reported in Table 3. BRL ratio: bankrupt loans to assets, RSL ratio: restructured loans to assets, Z-score $Z_{it} = \left(ROA_{it} + Capital\ ratio_{it}\right) / \sigma ROA_{it}$. The second row shows dependent variables, the third row shows the threshold values, the fourth row indicates low and high regimes, Mar: March, Sep: September, 00-15: 2000-2015, Obs: number of observations.

Table 5. Dynamic Panel Threshold Analysis for the Risk-Quantitative Easing Nexus (10-year Japanese government bond yield and Bank of Japan assets)

	1		2		3		4		5		6	
Dependent variable	BRL ratio		RSL ratio		lnZ-score		BRL ratio		RSL ratio		lnZ-score	
Threshold variable	Yield		Yield		Yield		BOJ assets		BOJ assets		BOJ assets	
Threshold estimates	1.032%		1.032%		1.484%		118,437,502 mil JPY		118,437,502 mil JPY		118,437,502 mil JPY	
95% confidence interval	[1.032% 1.032%]		[1.032% 1.032%]		[1.484% 1.484%]		[118,437,502 118,437,502]		[118,437,502 118,437,502]		[118,437,502 118,437,502]	
Impact of threshold variables	Est.	S.e.	Est.	S.e.	Est.	S.e.	Est.	S.e.	Est.	S.e.	Est.	S.e.
Low regime	-0.017***	0.0014	-0.0149***	0.0010	0.0856***	0.0113	-0.0347***	0.0098	-0.0286***	0.0050	-0.1219	0.0929
High regime	0.0131***	0.0049	0.0107***	0.0028	0.6513***	0.1882	0.0007	0.0015	-0.0010	0.0010	-0.1709***	0.0194
Intercept	-0.1427***	0.0253	-0.1222***	0.0151	-2.2627***	0.8002	0.6492***	0.1814	0.504***	0.0910	-0.9858	1.9967
Impact of covariates												
Boone	-0.1837***	0.0528	-0.122***	0.0296	0.7841***	0.2696	-0.1767***	0.0530	-0.1151***	0.0273	0.8562***	0.2814
Size	-0.0232***	0.0035	-0.0095***	0.0026	0.058**	0.0295	-0.0245***	0.0031	-0.0108***	0.0023	0.0543*	0.0283
Capital ratio	-0.4923***	0.1082	-0.0085	0.0390			-0.4981***	0.1034	-0.0142	0.0362		
Asset diversification	-0.1008***	0.0087	-0.0602***	0.0053	0.1431	0.1502	-0.1093***	0.0087	-0.0691***	0.0051	0.0609	0.1414
Liquidity	-0.0737***	0.0111	-0.0326***	0.0078	0.2455	0.2653	-0.0897***	0.0116	-0.0474***	0.0078	0.1494	0.2634
Revenue diversification	-0.0178***	0.0046	-0.016***	0.0035	0.1182*	0.0606	-0.0197***	0.0043	-0.0181***	0.0033	0.1033*	0.0562
GDP growth	0.0847***	0.0119	0.0712***	0.0078	0.3185***	0.0741	0.0081	0.0083	-0.0018	0.0053	-0.0463	0.0750
Market capitalisation	-0.0113***	0.0018	-0.0083***	0.0012	0.0479***	0.0082	0.0002	0.0011	0.0028***	0.0006	0.1032***	0.0067
Obs in low regime	1114		1114		2898		1156		1156		1156	
Obs in high regime	2377		2377		593		2335		2335		2335	

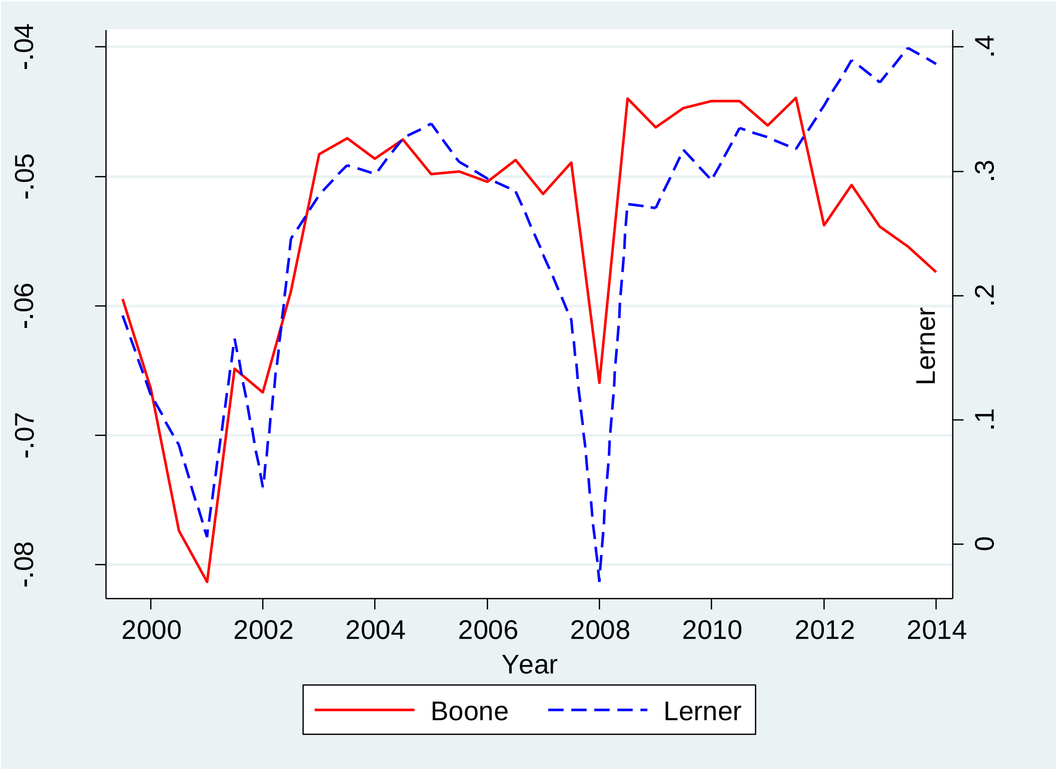
Notes: This Table reports the results from the dynamic threshold analysis using the first lag of the endogenous variable (Boone) as its instrument. The threshold variable is the 10-year Japanese government bond yield and Bank of Japan (BOJ) assets. BRL ratio: bankrupt loans to assets, RSL ratio: restructured loans to assets, Z-score $Z_{it} = \left(ROA_{it} + Capital\ ratio_{it}\right) / \sigma ROA_{it}$, lending rate=interest income on loans/loans and bills discounted, size=ln(total assets), capital ratio=equity/assets, asset diversification=securities/assets, liquidity=liquid assets/total assets, revenue diversification=non-interest incomes/operating income, market capitalisation is in natural logarithm. Est.: estimate, S.e.: standard error, Obs: number of observations. ***,**,*: significance at 1%, 5%, 10% level.

Table 6. Variance Decompositions.

	1				2				3			
Periods	Variables	QE	Boone	BRL ratio	Variables	QE	Boone	RSL ratio	Variables	QE	Boone	lnZ-score
5	QE	0.9997	0.0001	0.0002	QE	0.9997	0.0001	0.0002	QE	0.9888	0.0004	0.0108
5	Boone	0.0438	0.9561	0.0000	Boone	0.0292	0.9683	0.0025	Boone	0.0567	0.8303	0.1130
5	BRL ratio	0.2801	0.0765	0.6434	RSL ratio	0.1168	0.0166	0.8667	lnZ-score	0.0039	0.1429	0.8532
10	QE	0.9996	0.0001	0.0003	QE	0.9992	0.0001	0.0007	QE	0.9686	0.0012	0.0301
10	Boone	0.0835	0.9164	0.0000	Boone	0.0614	0.9350	0.0036	Boone	0.1030	0.7729	0.1240
10	BRL ratio	0.4679	0.0566	0.4756	RSL ratio	0.3089	0.0126	0.6785	lnZ-score	0.0068	0.1279	0.8653

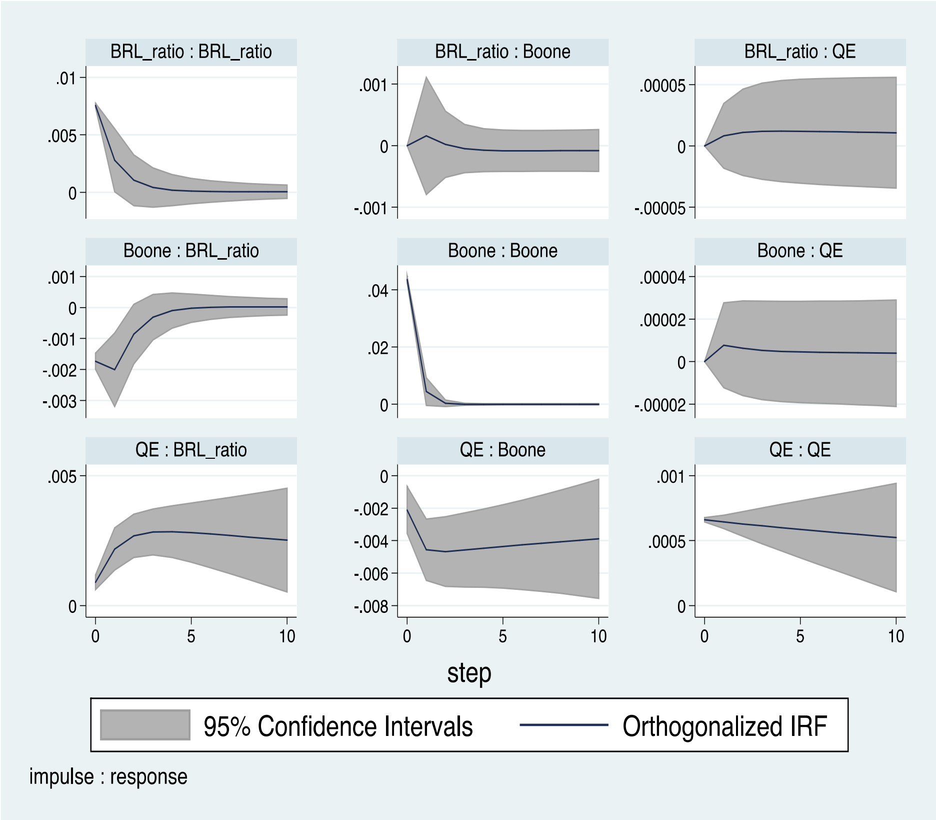
Notes: This Table reports the variance decompositions of the panel vector autoregression model for 5 and 10 periods ahead. There are 3 models, each has 3 variables: quantitative easing *QE* proxied by the lending rate, competition proxied by the Boone indicator, and risk. Column 1: risk is represented as bankrupt loan (BRL) ratio, column 2: risk is restructured loan (RSL) ratio, column 3: risk is taken as ln *Z*-score, $Z_{it} = \left(ROA_{it} + Capital\ ratio_{it}\right) / \sigma ROA_{it}$.

Figure 1. The Boone Indicator and the Lerner Index



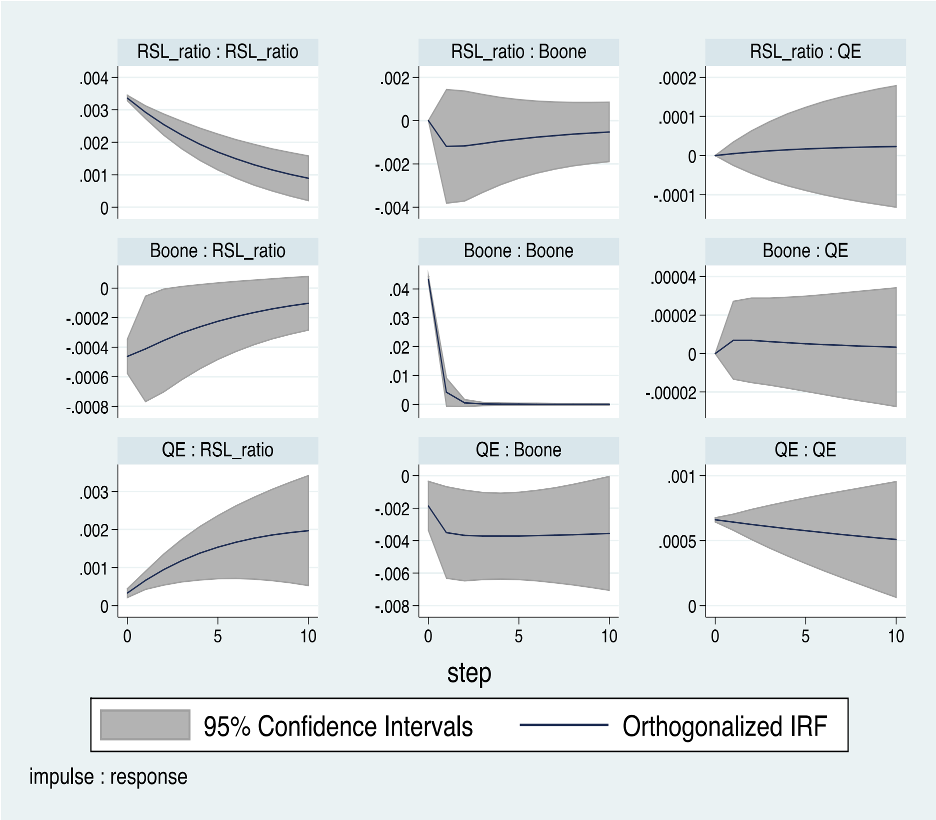
Notes: This Figure illustrates the average values of the Boone indicator and the Lerner index over time. Year denotes financial year.

Figure 2: Impulse Response Functions-Bankrupt loan ratio



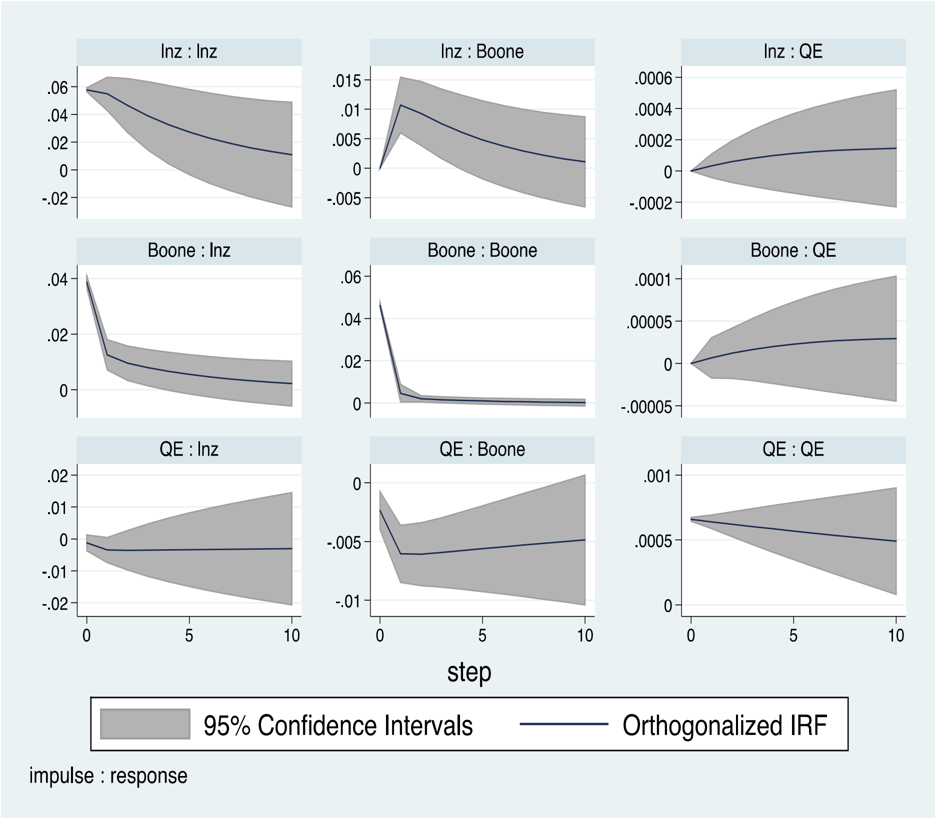
Notes: This figure illustrates the impulse-response functions (IRFs) of each endogenous variable with respect to one standard deviation shock in other variables. QE: Quantitative easing represented by bank lending rate; Boone is the Boone indicator of competition; BRL_ratio is bankrupt loan ratio; step: number of periods. Errors are 5% on each side generated by Monte-Carlo simulation.

Figure 3: Impulse Response Functions-Restructured loan ratio



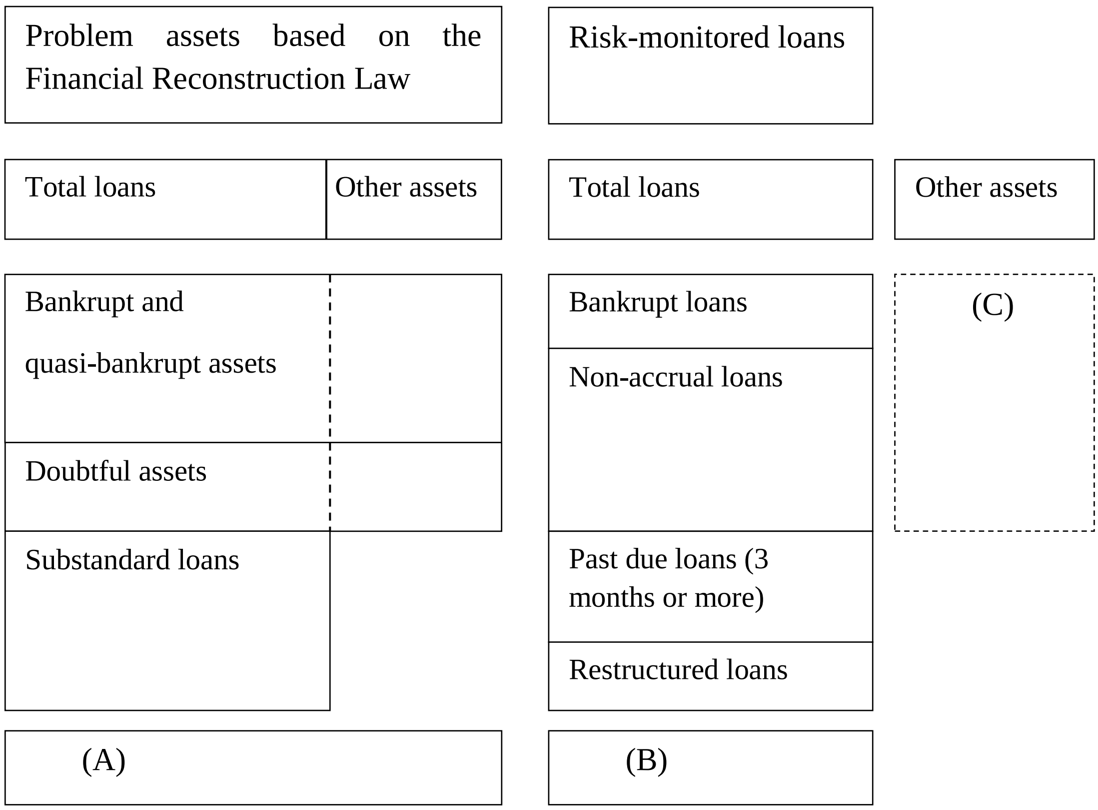
Notes: This figure illustrates the impulse-response functions (IRFs) of each endogenous variable with respect to one standard deviation shock in other variables. QE: Quantitative easing represented by bank lending rates; Boone is the Boone indicator of competition; RSL_ratio is restructured loan ratio; step: number of periods. Errors are 5% on each side generated by Monte-Carlo simulation.

Figure 4: Impulse Response Functions-Bank stability



Notes: This figure illustrates the impulse-response functions (IRFs) of each endogenous variable with respect to one standard deviation shock in other variables. QE: Quantitative easing represented by bank lending rate; Boone is the Boone indicator of competition; lnz is the natural logarithm of Z-score $Z_{it} = (ROA_{it} + Capital\ ratio_{it}) / \sigma ROA_{it}$; step: number of periods. Errors are 5% on each side generated by Monte-Carlo simulation.

Appendix A. Problem Assets Based on the Financial Reconstruction Law and Risk-Monitored Loans



Note: (A) – (B) = (C)

Notes: This Appendix presents the two classifications of problem assets in Japan. The difference between the two is other assets which are nonperforming assets (claims related to securities lending, foreign exchanges, accrued interests, suspense payments, customers' liabilities for acceptances and guarantees, and bank-guaranteed bonds sold through private placements). Risk-monitored loans are disclosed in accordance with the Banking Law. In this paper, Bankrupt loans are named after the sum of Bankrupt loans and Non-accrual loans; Restructured loans are named after the sum of past due loans over 3 months but less than 6 months and Restructured loans (The Japanese Bankers Association originally defined restructured loans as loans of which interest rates had been lowered. In 1997, the definition was extended to loans with any amended contract conditions and loans to corporations under ongoing reorganisation (Montgomery and Shimizutani, 2009)). Source: Interim report 2010-Sumitomo Mitsui Financial group.

Appendix B

B.1 The marginal cost

In order to attain values for the Boone indicator, we need to model bank marginal cost. In line with Fiordelisi and Mare (2014) and Fu et al. (2014), marginal cost is obtained from a flexible translog cost function specification[18]:

$$\ln TC_{it} = \alpha_0 + \alpha_1 \ln Q + \frac{1}{2}\alpha_2 \ln Q^2 + \sum_{j=1}^{2}\beta_j \ln P_j + \frac{1}{2}\sum_{j=1}^{2}\sum_{k=1}^{2}\delta_{jk}\ln P_j \ln P_k + \sum_{j=1}^{2}\gamma_j \ln Q \ln P_j$$
$$+\varphi_1 t + \frac{1}{2}\varphi_2 t^2 + \varphi_3 t \ln Q + \sum_{j=1}^{2}\varphi_j t \ln P_j + \varepsilon_{it} \tag{B1}$$

with total costs TC_{it}, total earning assets Q (loans, investments, and securities) (Delis, 2012), price of inputs P_j (which have to satisfy the condition of homogeneity of degree one), time trend t and a composed error term ε_{it}. Two input prices are incorporated: i) price of funds P_1 is defined as interest expenses divided by deposits and borrowed funds; ii) price of physical capital and labour P_2 as noninterest expenses divided by fixed assets[19].

The marginal cost *MC* for bank *i* at time *t* can be derived from equation (1) as follows:

$$MC_{it} = \frac{\partial TC_{it}}{\partial Q_{it}} = \frac{TC_{it}}{Q_{it}} \cdot \frac{\partial \ln TC_{it}}{\partial \ln Q_{it}} = \frac{TC_{it}}{Q_{it}}\left(\alpha_1 + \alpha_2 \ln Q + \sum_{j=1}^{2}\gamma_j \ln P_j + \varphi_3 t\right) \tag{B2}$$

B.2 The Boone indicator

The Boone indicator of competition has quite a few advantages in comparison with others. This measure accounts for both a lift in entry barriers or more aggressive interaction between market participants (Boone, 2008), while other indicators contain limitations or biases. As Beck (2008) argues, concentration ratios such as the Herfindahl-Hirschman index and three (five)-bank concentration ratio are rather unreliable measures of competition as they only weigh concentration levels. Concentration ratio could rise following an increase in competition, as uncompetitive participants would have to exit the market. Hence, if one

[18] Subscripts (*it*) are omitted for simplification.

[19] Due to data unavailability, we are unable to extract data from general and administrative expenses which include personnel expenses and non-personnel expenses associated to physical capital. Hence, we define the second input price in line with Fu et al. (2014).

interpreted higher concentration ratios as a proxy for uncompetitive markets, the results could be misleading (Schaeck & Cihák, 2014). Other measures of competition such as the Panzar-Rosse H-statistic and Lerner index also have some limitations. While H-statistic requires a priori assumption of long-run equilibrium operating markets (Panzar & Rosse, 1987), it is ambiguous whether the Lerner index captures the degree of product substitutability (Vives, 2008). Mirzaei and Moore (2014) argue that the H-statistic does not embrace the evolution of bank competition as there is only one score obtained over time. Even though time-varying scores are achievable (Bikker & Haaf, 2002; Jeon et al., 2011), they are either increasing or decreasing which may be inapplicable in effect.

Introduced by Boone et al. (2007) and Boone (2008), firms' (banks') market power can be measured through profit elasticity *β* in a simple profit equation:

$$\ln \pi_{it} = \alpha + \beta \ln mc_{it} + u_{it} \tag{B3}$$

where π_{it} and mc_{it} are profit and marginal costs of bank *i* at time *t*. *β* is the Boone indicator of market power which is expected to be negative as higher marginal costs would result in lower profits. Intuitively, in a competitive market, inefficient banks signified by comparatively high marginal costs are penalised more harshly since they will endure high loss in profits, compared to operating in an uncompetitive market. Hence, the larger the absolute value of *β*, the more intense the degree of competition.

In our paper, we employ the non-parametric methodology used in Delis (2012) to compute the Boone indicator for individual banks in each period. This allows us to create bank level estimates of competition. We estimate equation (3) by using a local regression analysis[20], which fits the relation between log profits and log marginal costs on the neighbourhood subsample of each observation to obtain individual β_{it}.[21]

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[20] According to Loader (1999), a local regression $Y_i = \mu(x_i) + \varepsilon_i$ with predictor variable *x* and response variable *Y* is estimated by smoothing the unknown function $\mu(x_i)$. This is obtained through fitting a polynomial model within a sliding window of *x*. Each point in the neighbourhood of *x* is assigned a weight corresponding to its distance from *x*. In particular, the closer the point to *x*, the larger its weight. The next step is to choose an optimal bandwidth *h* which controls the smoothness of fit and a smoothing window (*x*-*h*(*x*), *x*+*h*(*x*)). In other words, for each observation x_i, all neighbour points within the sliding window *h* are used in the following locally weighted least squares criterion: $\sum_{i=1}^{n} W\left(\frac{x_i - x}{h}\right)(Y_i - (a_0 + a_1(x_i - x)))^2$ with *W* is the weight function of

B.3 The Lerner index

We also use the Lerner index, another proxy of bank market power, to achieve a comprehensive analysis with different indicators of competition. The Lerner index is formulated as follows:

$$Lerner_{it} = \left(P_{Q_{it}} - MC_{it}\right) / P_{Q_{it}} \quad \text{(B4)}$$

where $P_{Q_{it}}$ is output price calculated as operating income divided by earning assets. This indicator captures pricing ability above marginal cost, which has been used extensively in the banking literature (Berger et al., 2009; Fiordelisi & Mare, 2014; Fu et al., 2014; Koetter et al., 2012). Values of the index are bounded between 0 and 1, with the former presenting perfect competition while the latter indicating pure monopoly. A negative Lerner index entails inability to price above marginal cost which might be a consequence of non-optimal behaviour (Fu et al., 2014).

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the form $W(u) = \begin{cases} \left(1-|u|^3\right)^3 & if\,|u|<1 \\ 0 & otherwise \end{cases}$ where $u = \left(x_i - x\right) / h(x)$. Following Delis (2012), we use the generalised cross-validation method to obtain our bandwidth of 0.42.

[21] In regression (B3), the Boone indicator is averaged over the entire sample across the whole examined period. Put differently, it cannot be measured for individual banks. To overcome this drawback, empirical research has modified this model to yield values of *β* for each period (Schaeck & Cihák, 2014; Van Leuvensteijn et al., 2011) by adding a time dummy and its interaction with marginal costs in order to increase the frequency of the indicator. However, the number of observations achieved from this approach does not rise significantly as they are average values for each period.

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